The Impact of the Parental Contribution on the Rate of Participation in Social Assistance: A Natural Experiment Approach[1].

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Abstract:

This study evaluates the impact of the parental contribution—introduced during the welfare reform of 1989—on rates of participation in social assistance. From a statistical perspective, this reform allows us to identify a control group comprising single individuals aged 30 and older. Our study groups consist of single persons under 30 subdivided into several age groups. Our empirical approach is based on "difference-in-difference" estimators often applied to natural experiments. This methodology can be generalized to account for other variables which may have differential impacts on the participation rates of the study and the control groups. Our results indicate that the parental contribution reduces the mean participation rate of the 20 and younger cohort by 19.4%. This impact falls to 12.1% for 21 year olds, and becomes insignificant for the over 21 group. In the 20 and younger cohort, the negative effect of the parental contribution considerably mitigates the positive effect of the benefits-scale parity introduced in 1989 for young claimants. Thus, thanks to the parental contribution, an expected increase of 22.0% in the participation rate of that cohort was limited to 3.0%.

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1. Introduction

The welfare reform of 1989 introduced several modifications to the way in which assistance is provided to underprivileged individuals and households. One of the principal changes was the introduction of benefits-schedule parity between childless individuals deemed fit to work under 30 years of age and older participants. Since the benefits for which they were eligible had previously been significantly lower than those given to over 30 claimants, this reform provided a substantial boost in their benefits schedule. For example, for those not participating in employability programs, the increase was 121%.

However, this reform also introduced a parental contribution (*Ministère de la Main-d'oeuvre et de la Sécurité du Revenu*, 1989, p. 11). This provision is based on a fundamental principle of the Civil Code (article 585), which stipulates that under normal conditions parents are financially responsible for their children. A financial contribution is required from those parents whose resources are deemed adequate, and it is factored into the calculation of the benefits payable to their adult children. This contribution, based on the parents' net income, may reduce or even void the child's eligibility for assistance. In this latter case the claimant is simply not admissible to the income security program. The parental contribution also makes it possible to better harmonize this assistance of last resort with the student loans and bursaries program. This latter program includes a parental contribution in the calculation of the level of loans and bursaries the student may receive (*Ministère de la Main-d'oeuvre et de la Sécurité du Revenu*, 1989, p. 12).

In terms of its distributive impact, a potential benefit of the parental contribution is to increase equity between youths from wealthier and poorer backgrounds. This is true to the extent to which the former benefit from parental assistance (when it is not declared) to which the latter do not have access.

By rendering the admissibility criteria more severe for certain individuals, the parental contribution has the direct effect of decreasing their rate of dependency, i.e. the amount of time they spend on the program. It may reduce either the frequency or the duration of their spells on social assistance. Moreover, it makes social assistance less appealing for some individuals, and thus pushes them toward a greater autonomy vis-à-vis the government. This behavioural effect compounds the direct impact of the parental contribution on the rates of welfare dependency. Finally, because of these mechanisms it is possible that the increased attractiveness of the welfare system created by the enhancements to the benefits schedule in 1989 may be largely annulled by the provision for a parental contribution in the case of very young adults.

What impact does the parental contribution have on the level of welfare dependence among these individuals? Can this impact be isolated from the other factors influencing this dependence (the benefits schedule, the business cycle, restrictions in employment insurance, changes in the minimum wage, etc.)? The purpose of this study is to answer these questions. More precisely, our goal is to evaluate the impact of the parental contribution on rates of participation in social assistance—in other words, on the proportion of APTE[4] claimants in the age groups affected by the parental contribution.

From a statistical perspective, we exploit those characteristics of the 1989 reform making it a *quasi-natural* experiment. Since the benefit schedule applying to single persons aged 30 and older was not directly affected by the reform, and since members of this group are furthermore potentially liable to have parental contribution claims laid against them, they comprise our control group[5]. Conversely, single adults under 30 years of age comprise the study group. Indeed, our model allows us to differentiate between several study groups according to age. This distinction is particularly useful since we would expect, for example, that the parental contribution would have a much greater impact on youths in the 18–21 cohorts than on those aged 22–25.

The econometric technique we use is based on “difference-in-difference” estimators often applied to the analysis of the results of natural experiments. In the context of regression analysis, the estimators are calculated as follows. The dependent variable is the rate of participation in social assistance by age group and by period. We first introduce the usual control variables into the model (the unemployment rate, the benefits schedule, the generosity of unemployment insurance, the minimum wage, seasonal indices, etc.). Notice that inclusion of the benefits schedule allows us to account for the impact of the provision for benefit-schedule parity in the reform. Next, we add three sets of explanatory variables into the model. The first set includes

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[4] APTE—Actions positives pour le travail et l'emploi. This is the name of the Quebec social assistance program for those who are able to work.

[5] PBK: Notice that it says further in the text that welfare recipient parents are not subject to the parental contribution for their adult children. Also, this would appear to undermine their usefulness as a control group, since they are now also directly affected by the parental contribution provision.

indices identifying each of the study groups. These variables capture the "fixed effects" by study group. Second, a dummy variable is added to reflect whether the observation occurs before or after the 1989 reform. This variable captures the post-reform impact common to all groups (control and study). Finally, we introduce terms for the interaction between this post-reform variable and each of the study-group specific index variables. These variables describe the post-reform impact on each study group, *above and beyond that captured by the common impact variable.* Since the benefit-schedule adjustment is already accounted for, the coefficients of these variables provide a natural measure of the effect of the parental contribution on each of the study groups.

The fact that the reform of 1989 contained several other provisions poses certain statistical problems, however. Thus, this reform led to the replacement of the SUPRET (Supplément au revenu de travail) program with the APPORT program (Aide aux parents pour le retour au travail). The purpose of these two programs is to supplement work income and to increase the financial incentives to finding employment. Now, while single individuals aged 30 or more were eligible for the SUPRET program, they are not for APPORT. Consequently, imputing effects to parental contribution that are in fact attributable to the substitution of the APPORT program for the SUPRET program may bias the results. Nonetheless, we do not believe that this impact is significant, given that SUPRET was not widely known, very complicated, and posed serious operational difficulties. For these reasons it was rarely used.

Another provision of the 1989 reform that must be considered is the expansion of access to employability-development programs already introduced by the government in 1984. In this analysis we shall assume that the index variable common to the study groups and the control group subsumes this provision. Finally, the reform also made provision to reduce premiums payable to claimants living in shared accommodations. The benefits-schedule variable partially captures this. Moreover, all other unobservable residual effects of this reform are assumed explained by the post-reform dummy variable in interaction with the index variable for the study group of claimants aged 25 to 29 years. This assumption is valid to the extent that this group is not affected by the parental contribution. Thus, the impact of the parental contribution on the 18–24 cohorts is measured by the interaction term between the post-reform dummy and the index variable for each sub-group, net of the residual effect of shared accommodations.

It is possible that estimates of the impact of the parental contribution are vulnerable to certain biases, to the extent that certain unobservable post-reform variables may impact the participation rates of the control and the study groups differently. It may be quite misleading to ascribe to the parental contribution some share of the impact which is in fact attributable to these variables. We account for these biases by introducing demographic variables, such as the weight of each cohort in the total population of single people 18–65 years old, as well as variables for the interaction between the observation years and the indices for the different study groups.

Section 2 summarizes the principles and the overall operation of the parental contribution, and profiles the impact of the program. In section 3 the statistical model is explained. Section 4 examines our data, while section 5 provides an analysis of the results. The last section presents the main conclusions of our study.[6]

2. The Parental Contribution

The parental contribution determines both the eligibility and the level of benefits payable to single adults considered "dependent" on their parents. An adult is considered dependent and subject to the parental contribution if none of the following conditions are met: having for a period of at least two years, not counting time during which he was a full time student, been self sufficient and lived outside of his father's or his mother's place of residence; having for at least two years held a full time paid job or having received employment benefits for such a job under to the provisions of Bill 1971; being, or having been, married; living common law with another person and having cohabited at some time with that person for a period of at least one year; having a dependent child; possessing a bachelor's degree; or being pregnant for at least 20 weeks, which needs to be attested by a medical certificate (*Loi sur la sécurité du revenu*).

Furthermore, an adult may be exempt from the parental contribution provision under certain circumstances: the parents are themselves welfare recipients, the parents cannot be located or are deceased, the parents refuse to provide the required information or to contribute financially. Finally, this provision is applicable for a maximum of three years. It begins when the claimant subject to the parental contribution either receives his first benefit or would have had he not been disqualified by his parent's income.

The monthly parental contribution amounts to 40% of the difference between the parents' income and various regulatory exemptions, divided by twelve. For purposes of these calculations the parents' income includes the amount indicated under "net income" in the tax declaration and a set of other incomes established by regulation.

The nature of these criteria is such that the program will tend to play a greater role in the case of young adults. For example, older individuals will have a higher probability of having lived outside of the parents' residence for at least two consecutive years. Moreover, if they are refused benefits at the age of 18 years, they will be eligible three years later without being subject to the parental contribution.

Table 1 presents the proportion, by age group, of single welfare recipients who were subject to the parental contribution in November, 1998. It also indicates who was not subject to that measure, and the reasons for their exemption.

Table 1 – Reasons for exemption from parental contribution for single beneficiaries of social assistance, by age group, November 1998

Reasons for exemption	**≤ 20 years**	**21–24 years**	**25–29 years**	**30 years**	**All single claimants**
	%				
1. Completion of 3 year wait or 2 years living separately and autonomously or 2 years working	14.8	45.4	60.4	41.0	41.4
2. Exempt adult (financial support and accomodations)	15.6	20.5	23.8	40.0	35.8
3. Pregnant for more than 20 months	2.5	0.8	0.2	-	0.2
4. Parents receiving welfare	34.5	16.6	3.4	0.1	3.6
5. Refusal to contribute	3.2	1.3	0.1	-	0.3
6. Parents unlocatable	3.1	3.6	2.4	0.2	0.8
7. Refusal to provide information	2.0	1.7	0.3	-	0.3
8. Other reasons for exemption	2.5	5.5	8.9	18.7	15.9
Subject to parental contribution	21.8	4.6	0.5	-	1.7
Total (%)	100.0	100.0	100.0	100.0	100.0
Total number	**15775**	**19609**	**22374**	**212875**	**270633**

Source: *Direction de la recherche, de l'évaluation et de la statistique, Ministère de la Solidarité sociale*, Quebec

[6] PBK: I corrected the contents of this paragraph.

The numbers reveal that 21.8% of single people aged 20 or below were subject to the parental contribution. This falls to 4.6% for the 21–24 year old group, and to 0.5% for the 25–29 cohort. As expected, this decline is largely explained by the expiration of the maximum period of three years during which the measure may be applied, or by the fact that the recipient lived more that two years away from the parents' residence or spent that long on the labour market. Thus, in the case of 21–24 year olds, 45.4% of claimants were exempt on the basis of these criteria. This percentage rises to 60.4% for the 25 to 29 cohort. Indeed, the vast majority of single people subject to the parental contribution are less than 21 years old. As the numbers in table 2 reveal, 76.9% of single people subject to the parental contribution were 20 or younger.

Table 2 – Number of recipients subject to the parental contribution by age, and impact on benefits paid, November 1998.

Cohort	No loss of benefits		With loss of benefits		Total	
	Number	%	Number	%	Number	%
≤ 20 years	1967	77.7	1471	75.9	3438	76.9
21–24 years	497	19.6	402	20.7	899	20.1
25–29 years	60	2.4	56	2.9	116	2.6
>30 years	7	0.3	9	0.5	16	0.4
Total	2531	100.0	1938	100.0	4469	100.0
%	56.6		43.4		100.0	

Source: *Direction de la recherche, de l'évaluation et de la statistique, Ministère de la Solidarité sociale*, Quebec.

Nonetheless, as we see in table 2, the parental contribution does not always have an impact on the level of benefits received by claimants subject to this measure. Of the total number of those affected, more that half (56.6%) suffered no loss of benefits in November, 1998. In addition, the proportion of beneficiaries subject to the parental contribution who incurred no loss of benefits has been increasing in recent years (table 3). In 1992, half of them had their benefits reduced. This proportion gradually fell in the following years to reach 44.1% in 1995. The exact reasons for this decline are currently unknown, but we may assume that the evolution of households' disposable income played a roll, since it is a direct contributor to the calculation of benefits.

Table 3 – Evolution of the number of claimants subject to the parental contribution, by the impact on benefits

Year	No loss of benefits		With loss of benefits		Total	Disposable Income (Variation %)
	Number	% of total	Number	% of total		
1990	2125	51.0	2043	49.0	4168	6.5
1991	2327	49.4	2384	50.6	4711	4.9
1992	2949	50.2	2930	49.8	5878	2.9
1993	3541	53.3	3106	46.7	6647	2.6
1994	3484	54.7	2882	45.3	6366	-0.3
1995	3431	55.9	2712	44.1	6142	4.9

Source: *Direction de la recherche, de l'évaluation et de la statistique, Ministère de la Solidarité sociale*, Quebec. *Comptes économiques du Québec, troisième trimestre* 1999.

In addition to having their benefits reduced, individuals subject to the parental contribution may in fact be ineligible to receive welfare, either because their application is rejected or because their benefits are clawed back. According to estimates of the *Ministère de la Solidarité Sociale*, the 20 and younger cohorts are most subject to denial of benefits or clawbacks (table 4), confirming the greater impact of the parental contribution on this group. The number of beneficiaries thus affected by this provision is much smaller for the 21 to 29 aged group, and virtually nil for those 30 or older. Over the period of the analysis, the number of applications rejected or whose benefits were clawed back is slightly higher for men (1834) than for women (1524).

Table 4 - Estimate of the number of applications rejected or clawed back because of the parental contribution, November 1997 – October, 1998

Cohort	Number	%
20 years and younger	2342	69,7
21 to 24 years	915	27,2
25 to 29 years	91	2,7
30 years or plus	10	0,3
Sex		
Women	1524	46,8
Men	1834	53,2
Total	**3358**	**100,0**

Source: *Direction de la recherche, de l'évaluation et de la statistique, Ministère de la Solidarité sociale*, Quebec.

This data allows us to evaluate the number and the relative preponderance of individuals directly affected by the parental contribution, in that either their application was rejected or their benefits were partially or entirely clawed back. In the subsequent analysis, we assume that the number of beneficiaries suffering a loss of benefits in November 1998 is representative of the situation throughout 1998. We also assume that the total number of rejections or clawbacks between November 1997 and October 1998 is indicative of the corresponding number for 1998. These two effects in conjunction affect 3813 individuals in the 20 and younger group (table 5), that is 27.1% of all claimants in that cohort. This impact falls to 1317 for the 21–24 cohort, or 7.7%, and proves virtually non extant (0.7%) for the 25–29 cohort.

Table 5 – Direct impact of the parental contribution

Cohort	Number of claimants rejected or clawed back	% of total claimants in 1998	Total claimants in 1998
20 years and younger	3813	27.1	14,097
21 to 24 years	1317	7.7	16,997
25 to 29 years	147	0.7	18,905

Source : *Direction de la recherche, de l'évaluation et de la statistique et DAFPG, Ministère de la solidarité sociale*, Quebec.

The very cursory overview we have just presented is based exclusively on information gleaned from the administrative records of the *Ministère de la solidarité sociale*. It does not account for any disincentive effect on participation rates, which can only be estimated using statistical techniques. The next section presents the statistical methodology used in this study to evaluate the global impact of the parental contribution on participation rates.

3. The Statistical Model

The introduction of the parental contribution creates two types of impact on welfare clientele: a direct effect and a behavioural effect. As we illustrated in section 2, the direct effect arises from the fact that, with no behavioural adjustment, some applicants who are subject to the parental contribution will be excluded from the program. This will be the case when incorporation of the parental contribution will push the claimant's income above the threshold of eligibility for welfare benefits. This factor will thus have the effect of reducing the number of entrants into the social security system while increasing the number of departures. The latter event will occur when the parental income of some claimants reaches a level at which they are no longer eligible for welfare.

The parental contribution may also have a behavioural impact on entry and withdrawal from welfare. Thus, because of the reduced level of the benefits, the parental contribution may incite the potential claimant to pursue his studies, to enter (or remain on) the labour market, or to live with a spouse without claiming benefits, rather than to become a welfare recipient. Moreover, by being financially independent while living away from his parents' home, the individual may hope to be exempt from the parental contribution in the future (after two years). In the short run, these considerations may incite him to stay off welfare. Also, the reduction in benefits makes the labour market relatively more attractive for a welfare recipient, which may lead him to withdraw from the system more rapidly. Similarly, an individual may be motivated to return to school when subject to the parental contribution and the loss of benefits.

Using appropriate econometric techniques it is possible to evaluate the global impact (i.e. the sum of the direct and behavioural impacts) of the parental contribution on entry into, and withdrawal from, the welfare system. Two approaches suggest themselves. The first is based on microdata (Fortin and Lacroix, [1997]) while the second uses semi-aggregated data. This latter technique is the one chosen for the study presented here. While this methodology is vulnerable to an aggregation bias, it is less subject to measurement errors than the microdata approach.

The dependent variable we seek to explain is the rate of participation in the welfare system, i.e. for each age group the ratio of the clientele of single individuals in the APTE program to the corresponding population. This variable combines both the direct and behavioural impacts of the parental contribution on entry into and withdrawal from social assistance. Variations in the welfare program's clientele are, of course, nothing other than the difference between entries and withdrawals for a given period.

To present the methodological approach of our analysis, it may be of some use to begin by presenting a very simplified version of the model. We shall first concentrate on the global impact of all the provisions of the reform (the benefit-schedule parity, the parental contribution, etc.). We then examine how several extensions may make our model more realistic and allow us to isolate the impact of the parental contribution. The basic idea underlying our approach exploits the fact that our data includes a study group and a control group. As mentioned in the introduction, it is natural to postulate that single, able-bodied individuals under the age of 30 constitute the study group, while the control group consists of those 30 and over.

For the moment we ignore all observable variables that may globally impact participation rates over the sample period (the business cycle, the parameters of the employment insurance program, the benefits schedule, etc.) other than the reform. The simplest estimator of the global impact of the reform is given by the difference in the means of the participation rates (in logs) of the study group before and after its implementation. This estimator, however, is biased to the extent to which observed changes are not necessarily entirely due to the effects of the reform. A variety of factors, including those enumerated above, may have also affected the participation rate during the period of our observations. To account for this, we begin by assuming that participation rate changes observed in the control group for the period in question are attributable to those factors. We further assume these factors will have the same impact on the participation rates of the two groups (study and control). Subtracting the differences in the means of the participation rates from the previously described estimator allows us to eliminate the common factors. The estimator thus obtained is called the "difference-in-difference" estimator (e.g. Meyer, Viscusi and Durbin [1995] and Mullahy [1999]). We write this:

$$\tilde{\alpha} = (\bar{y}_{SA} - \bar{y}_{SB}) - (\bar{y}_{CA} - \bar{y}_{CB}), \qquad (1)$$

where $\bar{y}_{ij}$ represents the sample mean of the participation rates (in logs) of group i for the period j. The index S corresponds to the study group, and C to the control group. A indexes the period subsequent to introduction of the reform, and B the preceding period. The first term on the right-hand side of equation (1) is the initial estimator, while the second term captures the impact of factors common to both groups. The difference between these two expressions thus allows us to isolate the global impact of the reform.

We can give the "difference-in-difference" estimator a very useful interpretation in terms of a regression model with index variables. Given the following model:

$$y_{it} = \alpha_0 + \alpha_1 d_S + \alpha_2 d_A + \alpha_3 d_S d_A + \mu_{it}, \qquad (2)$$

Where d_S is an index variable equal to 1 for the study group and 0 otherwise, d_A is an index variable equal to 1 if the observation is from the post-reform period and 0 otherwise, and where the μ_{it} are normally distributed random variables with mean 0 and variance σ^2. Assume further that the index variables are not correlated with these random variables. In equation (2) the parameter α_0 represents the mean of the control-group participation rates before introduction of the reform. The parameter α_1 captures the impact on this mean of belonging to the study group rather than to the control group. Thus, this parameter corresponds to a time-specific fixed effect applied to the first group. α_2 represents the after-reform impact common to the control and study groups, also corresponding to a time-specific fixed effect. Finally α_3 captures the effect of the post-reform on the study group, above and beyond that common to the two groups. This last coefficient thus corresponds to the impact of the reform on the participation rates of the study group. It is easy to show that the least squares estimator of α_3 is equal to the "difference-in-difference" estimator, $\widetilde{\alpha}$. We have:

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\begin{aligned}
E(y_{ij} \mid d_S = 1,\ d_A = 1) &= \alpha_0 + \alpha_1 + \alpha_2 + \alpha_3 = m_{SA}, \\
E(y_{ij} \mid d_S = 1,\ d_A = 0) &= \alpha_0 + \alpha_1 = m_{SB}, \\
E(y_{ij} \mid d_S = 0,\ d_A = 1) &= \alpha_0 + \alpha_2 = m_{SA}, \\
E(y_{ij} \mid d_S = 0,\ d_A = 0) &= \alpha_0 = m_{CB},
\end{aligned}
$$

Subtracting these equations in pairs, and then subtracting their differences, we obtain:

$$[(\alpha_0 + \alpha_1 + \alpha_2 + \alpha_3) - (\alpha_0 - \alpha_1)] - [(\alpha_0 + \alpha_2) - \alpha_0] = \alpha_3 = (m_{SA} - m_{SB}) - (m_{CA} - m_{CB}). \tag{3}$$

Finally, replacing the means m_{ij} with the sample means in equation (3), we obtain the least squares estimator of α_3. This corresponds to the difference-in-difference of equation (1).

In order to isolate the impact of the parental contribution it is important to include in the model variables for the real benefits schedule. These variables, which may assume different values for different age groups, allow us to account for benefit-schedule parity and accommodation sharing at the time of the reform. The estimator we obtain is a consistent estimator of the impact of the parental contribution if no other shock affects the *relative* participation rates of the study and control groups over the period. Now, different variables will clearly have different impacts on the groups. If we neglect to account for this, we run the risk of imputing to the parental contribution variations in the clientele which may, in fact, be due to these other factors. First, certain policies, such as successive contractions in employment insurance and regulatory changes to the labour law (e.g. labour standards) may have impacted the various clienteles differently during this period. Second, the business cycle, especially the recession of 1990–93, will have affected all participants, but its impact was probably much greater on the study group (the under 30 cohort) than on the control group (those over 30). Furthermore, seasonal effects may influence the quarterly evolution of participation within a given year. Third, cohort-specific effects may have modified the labour market conditions for the various age groups differently. Thus, a relative fall in the size of the population of 18–24 year olds would increase their relative scarcity on the labour market, leading us to expect a fall in their welfare participation rate, *ceteris paribus*. More generally, structural shocks at the supply and the demand level on labour markets may have had differential impacts on the participation rates of the different age groups.

In addition, it is important to relax the assumption that the parental contribution had a constant effect across the different cohorts comprising the study group. Indeed, the previous section allowed us to underline that this measure had a much greater impact on the youngest clientele (e.g. the 18–21 cohorts) than on the older groups (e.g. the 25–29 cohort).

The regression model presented above may be generalized to account for these various elements. On one hand, it is a matter of introducing as many index variables (individual and for the interaction with the post-reform dummy) as the number of retained study groups. On the other hand, we add a vector of explanatory variables whose goal is to account for factors other than the reform that may influence participation rates. These variables may also be included in the model in interaction with the index variables of the study groups to allow their impact to vary between sub-groups. Under these conditions, equation (2) becomes:

$$y_{it} = \alpha_0 + \alpha_1' D_S + \alpha_2 d_A + \alpha_3' D_S d_A + \alpha_4' x + D_S' \mathrm{A}\, x + \mu_{it}, \tag{4}$$

where D_S is now a vector of index variables for the study sub-groups, x is a vector of explanatory variables (the real benefits schedule, parameters of employment insurance, the minimum wage, the unemployment rate, index variables for the year corresponding to the *t*-th quarter, etc.), α_1, α_3 and α_4 are vectors of parameters and *A* is a matrix of parameters.

In equation (4) the vector of parameters α_3 measures the impact of the parental contribution for each study sub-group. Specifically, it measures the proportion of the post-reform impact on each of them that is in excess of that common to both the study and the control group, given by α_2. We can thus describe α_3 as a generalized difference-in-difference estimator. The vector α_1

identifies the fixed effects by age group. The vector α_4 measures the effect of each explanatory variable x_j (in x) on the participation rate of the control group. The coefficient a_{ij}, representing the matrix A, corresponds to the impact of the explanatory variable x_j on the study subgroup i, beyond its effect on the control group.

So far, we have assumed that the random terms μ_{it} were homoskedastic (i.e. constant variance). It is, however, more plausible to assume that the variance of these terms differs between age groups. We use the estimator for the variance-covariance matrix proposed by Greene ([1997], p. 635). This estimator is robust to all forms of heterorskedasticity, and constitutes an extension to White's estimator [1980] in the case of a model with fixed effects. We must point out that this estimator does not account for serial correlation in the errors[7]. This does not appear to pose much of a problem, however. Tests for serial correlation presented later do not allow us to reject the hypothesis of no correlation between the errors for the different age groups.

Finally, notice that some explanatory variables have an impact on the participation rates spanning several quarters. This is notably true of the unemployment rate, the generosity of the employment insurance system, and the real minimum wage. Integrating the current and lagged effects of these variables into the vector of explanatory variables x, the model in equation (4) becomes a distributed lag model. Since this study is primarily interested in the long-term impacts of the explanatory variables, we use the approach proposed by Wickens and Breusch [1988]. They use simple algebraic transformations of the distributed lag model to directly obtain the long-term multipliers. The coefficients of the natural logarithms of the explanatory variables represent these long-term effects when the lagged effects are expressed as first differences of the log of x.

It should be borne in mind that the problem of spurious regression, well known in the framework of nonstationary time series, is also an issue in nonstationary panel data[8]. Kao (1999), for example, shows that the ordinary least-squares estimator with nonstochastic fixed effects applied

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[7] Arellano (1987) proposed a robust estimator of the error covariance matrix which accounts for both serial correlation and heteroskedasticity. However, the asymptotique properties of this estimator assume that the number of individuals in the pooled observations tends toward infinity, while the time-indexed number of observations on each individual remains fixed. In this study we have the opposite situation, and this estimator is consequently not appropriate.

to equation (5[9]), and constrained to a single explanatory variable, is consistent. However, the *t*-statistic diverges from the standard normal. Consequently, statistical inference from the parameters of this model on the basis of the standard tables is misguided.

Several approaches have been developed to test variables generated by panel data for nonstationarity and for the presence of long-term relationships (cointegration) between them. We use the Levin and Lin (1993) test of stationarity on the variable for participation rates, and the Dickey-Fuller test adjusted for time series on the explanatory variables whose values remain fixed across the age groups. Ordinary least-squares estimation of equation (4) must yield stationary residuals for μ_{it} in the presence of one or more nonstationary variables in order to allow the use of standard tables. If the residuals are stationary, there exists cointegration between the variables. Pedroni's (1999) test for cointegration is the most appropriate one available to test the null hypothesis of non-cointegration in an equation incorporating several types of heterogeneous effects. This test allows for the presence of fixed effects with deterministic properties unique to each group in the pooled data, and for up to seven regressors whose effects may also vary between the elements of the pooling. As we shall see later, estimation of equation (4) in its final form incorporates interactive effects between the cohort-specific fixed effects and time. Moreover, this equation contains a large number of regressors because of lags on the explanatory variables. Pedroni's asymptotic results do not apply to this functional form, and we must limit application of the test to a simplified form of this model in which the interactive effects and lags on the explanatory variables are omitted. Under these conditions the test yields a indicator of probable cointegration between the variables.

4. The Data

Observations on the model's dependent variables, the participation rate by age group and period, are obtained as follows. We compute the numerator from data on the number of single persons participating in the APTE program beginning with the third quarter of 1989. Before this date the data excluded single individuals unfit for work. This is monthly data present in the administrative records of the *Ministère de la solidarité sociale*. It is available from 1977 to 1998

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[8] Cf. Banergee [1999], Maddala and Wu [1999], Kao [1999] and Pedroni [1999].
[9] PBK: There does not seem to be an equation 5.

and recorded by nine age groups, one group per year for the 18 to 24 cohorts, one group for 25 to 29 cohort, and one group aged 30 and over. In order to reduce serial correlation in the error terms we use the quarterly average of the data. Pooling the 792 observations (9 groups × 88 quarters) of time series and cross-section data gives us our database and satisfies the methodological requirements of the model.

To calculate the participation rates, each observation of the pooled data is divided by the total population of age group *i* in quarter *t*. The demographic data is taken from surveys of Statistics Canada. They are available on an annual basis, and we create quarterly values by simple interpolation.

Figures 2.1 to 2.4 at the end of this paper illustrate the evolution of the participation rate by age group. The participation rate of the 18–20 cohort grew less than that of the other age groups after the reform. The 1990–93 recession combined with the increase in the benefits schedule seems to have had a strong impact on claimants over 21 years of age, while figure 2.1 reveals a much lower impact on those 18 to 20. It is possible that for this age group the increased payments and the recession were of secondary importance compared to the parental contribution. Furthermore, very young adults may have an incentive to prolong their studies, or return to them, in an economic downturn. This phenomenon also has the effect of mitigating the impact of an economic slowdown on the participation rate of this group. Only an econometric study can isolate the role played by the parental contribution.

The database also includes pooled explanatory variables. These variables explain variations in the participation rates which are not attributable to the reform of 1989. The semi-aggregate nature of the data means that the pooling primarily includes variables which vary with time but remained fixed across groups[10]. All continuous variables are expressed as logs, and so their coefficients are elasticities. The real benefit schedules of the income security program are introduced into some versions of the model so as to account for elements of the 1989 reform other than the parental contribution. In addition, variables for the generosity of employment

[10] In the case of a panel of individuals observed during a period, we may have several types of explanatory variables: fixed for all individuals and variable with time, variable across individuals and fixed with time, variable across individuals and time.

insurance[11] and the real minimum wage are added to the model in order to account for other policies which may impact on the participation rate. The business cycle has a significant impact on the cyclical evolution of welfare claims. Its impact is measured by the unemployment rate of men 25 years old and over. Moreover, an index variable interacts with the unemployment rate as of the first quarter of 1990. The coefficient of this interaction variable accounts for the differentiated effect of the major recession of 1990–93 and the slow recovery that followed[12].

Furthermore, three variables indexing quarters are included to account for seasonal effects (the first three months of the year constitute the reference period). Cohort-specific effects are approximated by a population rate variable that measures the ratio of the population of a given age group to the total population aged 18 to 64. Finally, an index variable for the year, as well as variables capturing its interaction with the different age groups, are introduced to account for other unobservable influences on the participation rate.

Table 6 presents several sample statistics on the explanatory variables for the pre- and post-reform periods. The real benefits-schedule variables provide an approximate measure of the level of benefits actually paid to claimants (in 1992 dollars). Before 1989, the real benefits schedules are calculated from the schedule of benefits payable to single individuals for each age group. After 1989, a weighted average of the benefits schedule (including the Quebec sales tax[13]) incorporating categories of employability and accommodation sharing[14] is computed[15][16]. The schedule parity created in 1989 for under 30 claimants (combined with other sources of

[11] The generosity of employment insurance is measured as the ratio of the maximum number of weeks of benefits to the minimum number of weeks of work required to qualify for benefits. This ratio evolves as a function of a moving average of three months of the unemployment rate (not seasonally adjusted). The net rate at which employment insurance replaces earnings was also introduced, but was never significant.

[12] The unemployment rate, which is already included in the model, captures the mean impact of the business cycle on participation rates. The magnitude and duration of business cycles vary over from one to the next, and they may have nonlinear effects, which is why we incorporate an variable capturing the interaction of unemployment with the quarters from 1990–98.

[13] PBK: The relevance of the Quebec sales tax may require some explanation to those unfamiliar with the system.

[14] The accommodation sharing provision was introduced into the *Loi sur la sécurité du revenu* on August 1, 1989. It essentially reduces by $100 the allocation to a claimant who lives in a shared accommodation.

[15] Construction of the aggregate benefits schedule occurs in two steps. First, a weighted mean (M1) of the income security benefits schedule is calculated. The weights are a mean of the post-reform monthly averages of the number of single claimants of APTE subject to one of the four schedules of employability and the total number of single APTE claimants. Second, the amounts ($85, $93, or $100 according to the period) from the accommodation sharing provisions are deducted from M1 to yield M2. A weighted average of M1 and M2 is derived using weights computed from the post-reform average of the ratios of single APTE recipients sharing accommodations (or not) to the total number of APTE recipients given by the aggregate schedule.

[16] PBK: I did my best to translate the previous footnote accurately, but the calculations are not clear to me.

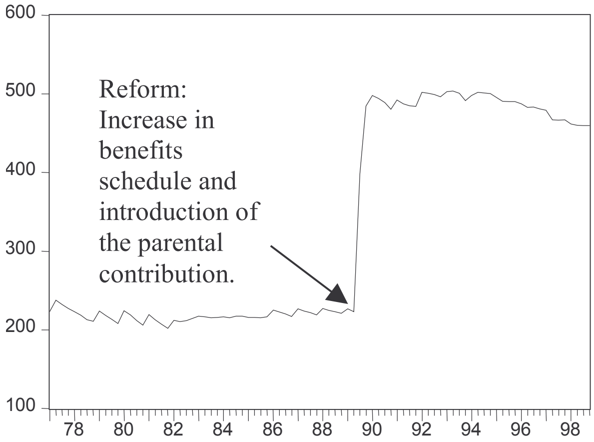
variation in the schedules) caused a 122% increase in real benefits paid during the post-reform period to this clientele. Naturally, this stimulated an influx of youth participation in the program. Figure 1 illustrates how little the schedules fluctuated before and after the reform of 1989. This feature does not facilitate our task of identifying the impact of the parental contribution, owing to the strong collinearity between the variable for the real benefits schedule and that measuring the parental contribution.

Table 6 – Sample characteristics, observed mean of the influences during the pre- and post-reform periods

Variable	Pre-reform	Post-reform	Change (%)
Real benefits schedule, < 30 years	218.3 (6.7)	485.4 (25.3)	122.4
Real benefits schedule, ≥ 30 years	590.6 (13.1)	488.6 (21.9)	-17.3
Unemployment rate	9.1 (1.6)	10.4 (1.4)	1.3
Real minimum wage	6.3 (0.9)	5.9 (0.4)	-6.3
Nonpecuniary generosity of E.I.	3.1 (0.5)	2.1 (0.6)	-32.3
Population rate (%)			
18 years	2.8 (0.4)	2.1 (0.1)	-0.7
19 years	2.9 (0.4)	2.1 (0.1)	-0.8
20 years	3.0 (0.4)	2.1 (0.1)	-0.9
21 years	3.0 (0.3)	2.1 (0.1)	-0.9
22 years	3.1 (0.2)	2.1 (0.1)	-1.0
23 years	3.1 (0.1)	2.1 (0.2)	-1.0
24 years	3.0 (0.1)	2.2 (0.2)	-0.8
25–29 years	14.7 (0.1)	12.2 (1.5)	-2.5
≥ 30 years	64.4 (2.0)	73.1 (2.0)	8.7
Number of observations	50	38	

Notes: Standard errors in parenthesis. The statistics are calculated for the periods 77.1 to 89.2 and 89.3 to 98.4 for the pre- and post-reform period respectively.

Figure 1 — Real benefits schedule for under-30 beneficiaries



The numbers in table 6 reveal little difference between the mean unemployment rates over the pre- and post-reform periods. The 1990–93 recession had significant repercussions on the labour market, however, and consequently on the rates of participation in social assistance.

The mean real minimum wage does not differ significantly before and after the reform, but it does show steady growth as of 1986. Also, the substantial changes in the employment insurance program implemented during the 1990-s, especially the contractions in the number of weeks of admissibility to the program and changes in the number of weeks of work required to qualify, had the net effect of reducing its generosity. The index for the generosity of employment insurance provides a measure of the extent of these changes. Over the post-reform period, the average value of this index fell by 32%, occasioning the arrival of new welfare claimants and reducing incentives to leave the program. Thus, we must expect that the global impact of restrictions to the employment insurance program was to increase participation rates. Finally, the mean population rates, i.e. the ratio of the population of a specific cohort to the total population of 18 to 64 year olds, declined for all under 30 age groups over the course of the post-reform period. For individuals aged 30 or more, this rate increased significantly (by 8.7%).

5. Results

Table 7 presents the results of the difference-in-difference estimates for the simplified model, revealing the global impact of the reform (including the benefits-schedule parity and the parental contribution) on the participation rates by age group. This estimator is calculated from the various components of the right-hand side of equation (3).

According to the numbers in this table, mean participation rates fell between the pre- and post-reform periods for the under 20 cohort (column 5), but we see that the extent of this fall decreases with age. Conversely, participation rates grew for the 21 to 29 cohort, and this increased with age. However, these changes reflect the entire range of factors that may impact on the evolution of welfare clienteles. Thus, it is important to net out those contributors whose effect is common to the control and study groups. Column (6) in table 7 presents the percent change in the mean participation rate for the control group (age 30 and over), a 23.2% increase between the two periods. The difference-in-difference estimators reproduced in column (7) correspond to the variation in the mean participation rate for each study group, adjusted downward by the increase in the control group. According to our results, the 1989 reform boosted the participation rate among 23–29 year olds, and this impact is statistically significant. Furthermore, it also induced a statistically significant fall in this rate for the 21 and younger cohorts. This decline in the rates proves even more pronounced in the case of the youngest age groups.

Table 7 – Difference-in-difference estimators of the impact of the 1989 reform on participation rates (logged) by age

	Mean participation rate (logged) Study group, Singles, 18–29 years		Mean participation rate (logged) Control group, Singles, ≥ 30 years		Differences		Difference-in-difference
	Post-reform (1)	Pre-reform (2)	Post-reform (3)	Pre-reform (4)	[(1)-(2)] (5)	[(3)-(4)] (6)	[(5)-(6)] (7)
18 years:	-3.077	-2.813	-3.267	-3.499	-0.264	0.232	-0.496
	(0.159)	(0.403)	(0.073)	(0.391)	(0.055)	(0.069)	(0.078)
19 years:	-2.863	-2.694	-3.267	-3.499	-0.169	0.232	-0.401
	(0.174)	(0.416)	(0.073)	(0.391)	(0.057)	(0.069)	(0.080)
20 years:	-2.866	-2.858	-3.267	-3.499	-0.008	0.232	-0.240
	(0.198)	(0.418)	(0.073)	(0.391)	(0.058)	(0.069)	(0.082)
21 years:	-2.890	-3.053	-3.267	-3.499	0.163	0.232	-0.069
	(0.221)	(0.412)	(0.073)	(0.391)	(0.058)	(0.069)	(0.082)
22 years:	-2.923	-3.246	-3.267	-3.499	0.323	0.232	0.091
	(0.241)	(0.405)	(0.073)	(0.391)	(0.058)	(0.069)	(0.082)
23 years:	-2.968	-3.412	-3.267	-3.499	0.444	0.232	0.212
	(0.260)	(0.391)	(0.073)	(0.391)	(0.058)	(0.069)	(0.082)
24 years:	-3.013	-3.549	-3.267	-3.499	0.536	0.232	0.304
	(0.267)	(0.380)	(0.073)	(0.391)	(0.057)	(0.069)	(0.081)
25–29 years:	-3.140	-3.841	-3.267	-3.499	0.701	0.232	0.469
	(0.259)	(0.379)	(0.073)	(0.391)	(0.057)	(0.069)	(0.080)

Notes: Standard errors in parenthesis. The mean pre-reform participation rate is the average of the rate (in logs) for the period 77.1 to 89.2, while the post-reform rate is based on the period 89.3 to 98.4.

While these estimation results are probably tainted by significant biases, it remains true nonetheless that they corroborate the hypothesis that the parental contribution caused a decline in the participation rates in the case of the youngest clienteles. This is true since, to all appearances, the other major element of the reform, the benefit-schedule parity for the below 30 cohort, had the impact of increasing their participation.

As we mentioned in section 4, it is possible to broaden the approach so as to account for the observable and unobservable variables that may have influenced the participation rates of the study groups relative to the control group. It is simply a matter of developing a regression model including the relevant explanatory variables [cf. Equation (4)]. Table 8 at the end of this paper presents the estimation results for the participation rates when we sequentially introduce various explanatory variables into the model to better isolate the impact of the parental contribution.

Column (A) presents the estimators of equation (2) generalized to several study groups. Thus, it is a regression model exactly corresponding to the difference-in-difference estimators in table (7). The coefficients associated with the variables *D18Dp* to *D2529Dp* (age, post-reform) are thus estimators of the impact of the 1989 reform on the different age groups. They represent the proportion of the impact of the reform that exceeds the common impact captured by the coefficient of the variable *Dp*. Notice that the standard errors used to calculate the *t*-statistics are probably not consistent, since they impose homoskedasticity on the error terms across age groups. Consequently, we also present *z*-statistics obtained from the covariance matrix proposed by Greene ([1997], p. 635). These statistics are robust to heteroskedasticity and their asymptotic distribution is that of student's *t*.

Columns (B) to (F) progressively introduce the control variables into the model. Column (B) incorporates the quarterly variables (three variables per age group) to account for seasonal influences in the evolution of the participation rates. These variables have very little impact on the estimated coefficients—we cannot reject the hypothesis that the seasonality effects are nil (*F*-statistic = 0.174) at the five percent confidence level.

Column (C) adds the real benefits-schedule variables into the model (for the age 24 and younger cohorts, the 25–29 cohort, and the 30-plus cohort), a variable for the real minimum wage, a variable for the generosity of employment insurance, a population-rate variable (fraction of each group in the 18–65 aged population), and finally variables for the unemployment rate in interaction with the age groups. The lag structure retained for these variables is presented at the top of the table. It was chosen on the basis of several specifications and numerous tests. The reported coefficients for these variables correspond to their long-term effects on the participation rates. The Fisher test confirms (F=86.886) that these variables jointly contribute a significant explanatory power to the evolution of the participation rates. Individually, these variables are also all significant, except the benefits schedules for claimants under 30 years of age. Since this specification accounts for variations in the benefits schedules, the coefficients of the post-reform age variables are estimators of the impact of the parental contribution by age group. It is interesting to note that the effect is negative, and significant, for claimants under 21 years of age, and that it decreases with age. Moreover, for this age group, the effect proves much greater than the global impact of the reform as estimated by the corresponding coefficients in column (A). This result is not surprising, since the global impact of the reform includes the benefit-schedule parity effect. This, of course, works in the opposite direction of the parental contribution.

Column (D) introduces the annual fixed effects. The purpose of these variables is to account for the unobservable shocks which may also impact on the participation rates (in percentage) of the different age groups. The Fisher test (F=52.127) confirms the significant impact of these factors on the evolution of the rates. Furthermore, they have the effect of reducing the impact of the parental contribution in the case of the 18–21 cohort vis-à-vis the estimators in column (C). The estimated coefficients indicate that the parental contribution reduces by 78.0% and 69.9%, respectively, the mean participation rate of 18 and 19 year old youths, and by 54.0% that of the age 20 cohort. Conversely, according to tests performed using the *z*-statistic, the parental contribution has no significant impact on groups of older single individuals. Column (E) adds the variable for the interaction between the unemployment rate and the quarterly indices for the period 1990–98. The test statistic confirms that this variable is significant at the five percent level. Moreover, the parental contribution has a negative and significant impact on the participation rates of all age groups below 25.

The last column, column (F), in table 8 presents the results from the most complete specification. This allows the annual effects to vary between age groups. Thus, we aim to account for unobservable variables (e.g. structural changes in the labour market for each age group) that may influence the *relative* participation rates of the study and control groups over the sample period. The Fisher statistic allows us to conclude that the annual variables by age group are jointly significant. In addition, a Lagrange multiplier test proposed by Baltagi ([1995], p. 93) leads us to not reject the absence of serial correlation of the residuals in this specification (statistic ML = 0.0102). The coefficients of the post-reform age variables in this column thus yield an appropriate measure of the impact of the parental contribution. The post-reform age effects for the 20 and younger group are all significant, but weaker than in the previous specifications. Thus, our results indicate a reduction of about 19% in the participation rate attributable to the parental contribution. This decline is only to 12.1% for the age 21 cohort. Finally, this provision has no significant effect on the participation rates of older claimants. These results corroborate the descriptive statistics on the reductions, rejections and benefit clawbacks which indicate that 18 to 20 year old claimants are more affected by the parental contribution.

It is of some interest to take a closer look at the other coefficients in column (F), since this specification generates estimators that are more comprehensive than those presented in the other columns of the table. According to the results from this specification, a 10% increase in the benefits schedule effects an increase of 1.80% in the participation rate for the 24 and younger cohort, and of 2.91% for the 25–29 cohort. These results may be used to analyse the impact of the introduction of benefit-schedule parity in 1989. This policy increased the mean benefit schedule by 122.4%, which would augment the participation rate among those aged 24 and less by 22.0%, and that of the 25–29 cohort by 35.6%. According to these calculations, the 1989 reform would have had the net effect of increasing by 3.0% (=22.0%–19.0%) the participation rate among those aged 20 and less. Thus, the parental contribution virtually cancelled the impact of the benefit-schedule parity. Since the parental contribution had no significant effect on the over 20 cohort, the reform led to a substantial jump in the participation rate of this group relative to those 20 and younger.

According to the results of this specification, an increase in the minimum wage reduces the global participation rate, while a decrease in the generosity of employment insurance leads to a significant increase. Furthermore, our results confirm that variations in the unemployment rate among those aged 25 and above induce changes, in the same direction, in the participation rates. It is also noteworthy that younger claimants are much more sensitive to the business cycle.

The Fisher test confirms that the participation rates of men and women differ significantly[17]. The substantial impact of the parental contribution on the 18–20 cohort is higher among women (20.2%) than among men, for whom it reaches 16.5% on average [Column (F), tables 9 and 10]. The parental contribution also reduces by 21.1% the participation rates of women aged 21. Those between 22 and 24 years old experience a milder impact (12.1%) on their participation rates, but *z*-tests confirm that these results have a greater probability of error. As to men aged 21 or more, they are not affected by the parental contribution. In summary, this provision has a greater impact on women than on men.

Finally, the Dickey-Fuller test of stationarity does not permit us to reject the null hypothesis of nonstationarity in the unemployment rate, in the real minimum wage, in the nonpecuniary

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[17] We apply the Fisher test described by Baltagi ([1995], p. 49) and in Hsia ([1986], p. 15). The *F*-statistic is here equal to 824.932, far greater than the threshold for significance.

generosity of employment insurance, and in the real benefit schedules for the two cases of presence and absence of a deterministic trend in the estimated equation (table 11). Moreover, the Levin and Lin test also confirms that the pooling of the participation rates is nonstationary. To ensure the validity of the preceding statistical inferences, Pedroni's cointegration test is applied to the model with neither interactive effects between age and time nor the lagged explanatory variables, and this for the reasons previously evoked. As indicated in table 12, the test rejects the null hypothesis of the absence of cointegration (nonstationary residuals) for each of the three statistics calculated. Since the model estimated for the test is not identical to the functional form finally yielding the impact of the parental contribution [column (F)], the Pedroni test only indicates the probable presence of cointegration between the model's variables.

6. Conclusion

This study examines the impact of the parental contribution on participation rates in social assistance, i.e. on the proportion of different demographic groups benefiting from assistance of last resort. A descriptive analysis of the data reveals that this provision particularly reduces or eliminates the benefits of young claimants between 18 and 20 years of age. We also find that the parental contribution may constitute a disincentive to becoming a claimant, and even encourage beneficiaries to leave the program more rapidly. These direct and behavioural effects reduce the participation rate.

In order to measure the global impact of this provision, we use a statistical approach based on difference-in-difference estimators. Implementation of this technique requires the definition of control and study groups. In our analysis, the control group comprises single individuals aged 30 and over, while the study groups consist of those under 30 categorized into age groups. The approach essentially involves creating the most precise association possible between the percent changes in participation rates and the period following the introduction of the parental contribution[18]. Our analysis accounts for influences common to the study and the control groups that may have an impact on the evolution of participation rates. It also seeks to isolate the impact of the parental contribution from shocks which may contribute to changes in the relative

[18] PBK: this sentence is not clear.

participation rates of these groups (the business cycle, cohort-specific effects, changes to the benefit schedule, etc.).

According to the results of our estimation, the parental contribution reduces by 19.4%, on average, the participation rates of single individuals aged 20 or less, and by 12.1% those of single 21 year olds. The provision does not have a significant impact on older cohorts. Among those 20 and younger this negative impact of the parental contribution almost entirely cancels the positive impact of the benefit-schedule parity extended to those under 30 by the 1989 reforms. Owing to the parental contribution, the global effect of the reform was a 3.0%, rather than a 22.0%, increase in claimants aged 20 or less.

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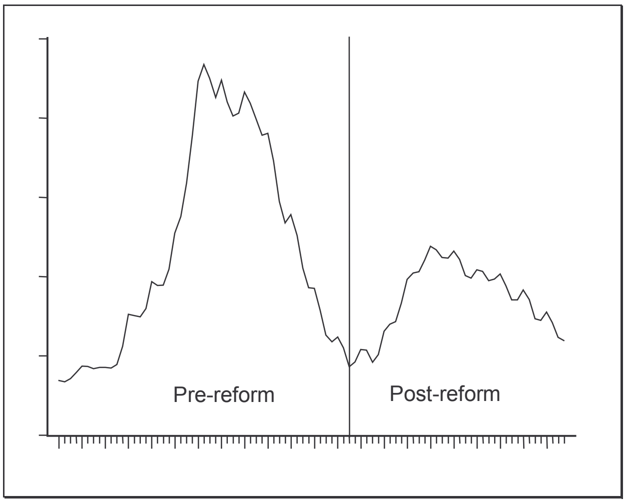
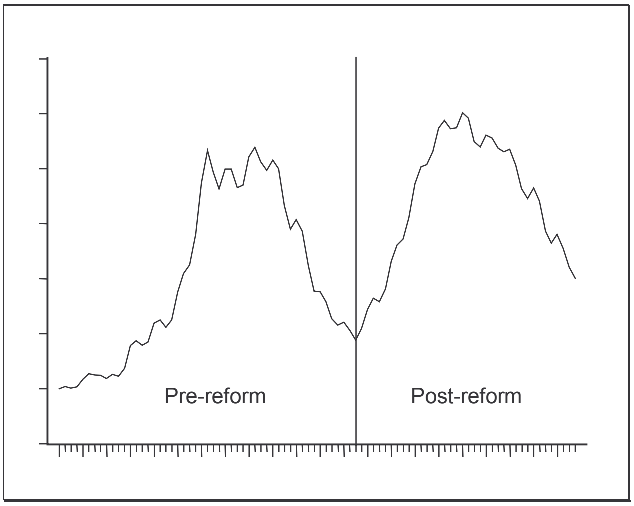
Welfare participation rates by age groups,
pe- and post-reform Periods

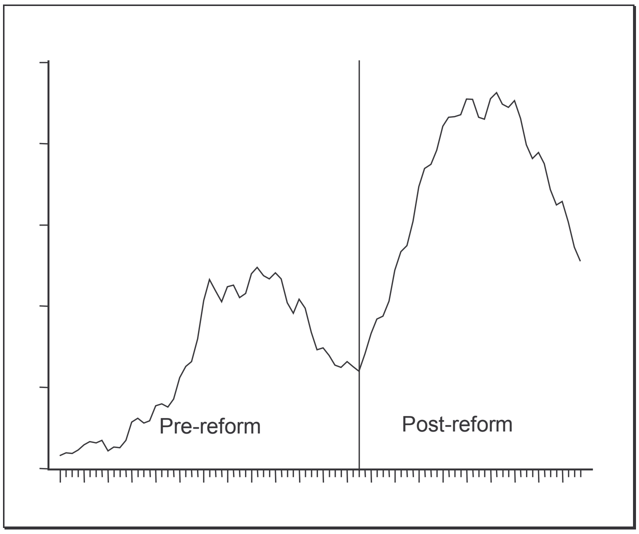
Figure 4.1
7. 18–20 cohorts



8. Figure 4.2
9. 21-24 cohorts



10. Figure 4.3
11. 25–29 cohort



12. Figure 4.4
13. 30 years or over cohort

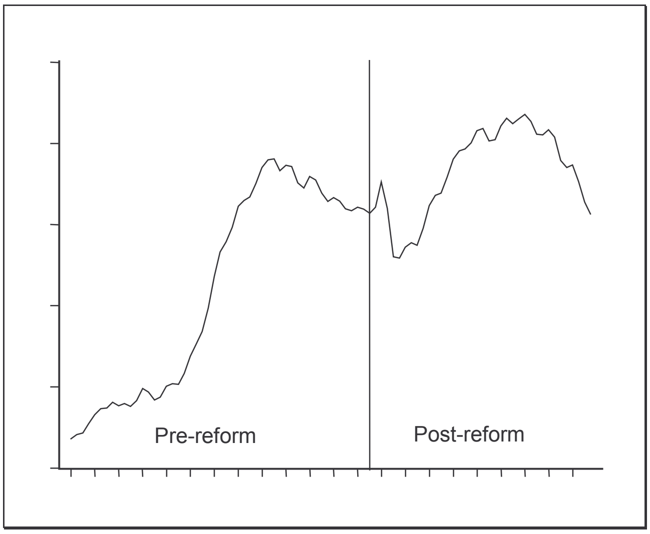


Table 8
Estimates of welfare participation rates

Difference-in-difference model, OLS, 77.1 to 98.4, quarterly data
Lag specification: *MW*= 1, *GEI* = 4, *U* =4. (*)

	(A)			(B)			(C)			(D)			(E)			(F)		
Fixed effects by age group	Yes			Yes			Yes			Yes			Yes			Yes		
Seasonal effects by age group	No			Yes			Yes			Yes			Yes			Yes		
Annual effects	No			No			No			Yes			Yes			Yes		
Annual effects by age group	No			No			No			No			No			Yes		
Effect of the 90-93 recession	No			No			No			No			Yes			Yes		
Effect of other factors (unemployment rate, benefits schedule, etc.)	No			No			Yes			Yes			Yes			Yes		
Fisher Test (confidence level = 5%). Each test is sequential: (B) with respect to (A), (C) with respect to (B), etc.				• F = 0,174 ➔ Seasonal effect not significant			• F =86,886 ➔ Exogenous factors significant.			• F =52,127 ➔ Time effect significant.			• F =4,401 ➔ Recession effect significant.			F =6.852 ➔ Age*time effect significant.		
	Coeff	***t*-ols**	***z*-asy (**)**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**
Constant	-3.499	-76.737	-118.146	-3.493	-46.280	-81.072	-10.043	-3.672	-7.434	-7.090	-3.556	-3.191	-4.910	-2.501	-2.783	-0.751	-0.239	-0.432
Common Post Effect (*Dp*)	0.231	3.336	7.272	0.232	3.295	7.313	0.077	0.887	1.929	0.249	3.623	2.754	0.317	4.582	6.094	0.010	0.122	0.256
Post 18 years effect (*D18Dp*)	-0.496	-5.056	-7.120	-0.496	-4.974	-7.153	-0.974	-6.018	-5.804	-0.780	-5.309	-3.360	-0.684	-4.782	-6.591	-0.192	-1.647	-3.185
Post 19 years effect (*D19Dp*)	-0.401	-4.084	-5.572	-0.400	-4.008	-5.584	-0.887	-5.495	-5.315	-0.699	-4.765	-3.027	-0.597	-4.172	-5.910	-0.194	-1.660	-3.278
Post 20 years effect (*D20Dp*)	-0.240	-2.445	-3.251	-0.238	-2.392	-3.250	-0.721	-4.477	-4.349	-0.540	-3.678	-2.341	-0.499	-3.489	-4.998	-0.195	-1.674	-3.334
Post 21 years effect (*D21Dp*)	-0.068	-0.690	-0.905	-0.066	-0.662	-0.889	-0.537	-3.338	-3.229	-0.360	-2.453	-1.563	-0.413	-2.885	-4.120	-0.121	-1.037	-2.024
Post 22 years effect (*D22Dp*)	0.092	0.934	1.212	0.094	0.940	1.249	-0.372	-2.312	-2.215	-0.196	-1.339	-0.851	-0.380	-2.651	-3.817	-0.108	-0.928	-1.761
Post 23 years effect (*D23Dp*)	0.213	2.172	2.813	0.215	2.157	2.857	-0.251	-1.561	-1.470	-0.074	-0.506	-0.320	-0.378	-2.634	-3.838	-0.101	-0.866	-1.541
Post 24 years effect (*D24Dp*)	0.304	3.103	4.039	0.306	3.073	4.093	-0.172	-1.068	-1.002	0.010	0.069	0.044	-0.383	-2.658	-3.966	-0.105	-0.899	-1.643
Post 25–29 years effect (*D2529Dp*)	0.469	4.783	6.307	0.471	4.725	6.372	0.017	0.060	0.037	0.069	0.329	0.185	0.056	0.288	0.330	-0.161	-0.833	-1.659
Real b s 18–24 years (*D1824Bm*)							0.373	2.430	2.168	0.074	0.513	0.329	0.158	1.120	1.506	0.180	1.877	2.894
Real b s 25–29 years (*D2529Bm*)							0.267	0.802	0.485	0.160	0.675	0.384	-0.292	-1.256	-1.310	0.291	1.208	2.253
Real b s 30 years + (*D30+Bm*)							1.129	2.600	5.303	1.037	3.524	3.072	0.961	3.300	3.800	0.034	0.070	0.136
Real minimum wage (*MW*)							-0.631	-11.613	-9.570	-1.008	-2.765	-3.125	-1.575	-4.330	-4.827	-1.634	-7.175	-7.931
EI generosity (*GEI*)							-0.416	-11.752	-12.351	-0.183	-1.584	-1.788	-0.339	-2.907	-3.147	-0.338	-4.662	-5.379
Population rate							0.563	8.459	6.456	0.392	8.210	5.685	0.666	13.374	10.369	0.264	1.137	1.045
Unemploy rate 18 years (*D18u*)							1.956	19.994	18.705	0.954	7.325	6.670	0.987	8.097	7.445	1.055	6.141	4.134
Unemploy rate 19 years (*D19u*)							2.096	21.601	24.510	1.105	8.492	8.893	1.118	9.188	10.041	0.996	5.782	5.451
Unemploy rate 20 years (*D20u*)							2.140	22.201	27.014	1.160	8.924	9.967	1.127	9.278	10.567	1.014	5.903	4.527
Unemploy rate 21 years (*D21u*)							2.119	22.082	27.587	1.151	8.848	10.212	1.066	8.773	10.196	0.936	5.453	5.374
Unemploy rate 22 years (*D22u*)							2.099	21.905	25.182	1.136	8.733	10.011	0.996	8.188	9.394	0.791	4.607	5.222
Unemploy rate 23 years (*D23u*)							2.058	21.488	21.261	1.098	8.439	9.406	0.915	7.524	8.449	0.801	4.663	4.745
Unemploy rate 24 years (*D24u*)							2.007	20.965	18.931	1.049	8.065	8.613	0.839	6.889	7.579	0.739	4.307	6.258
Unemploy rate 25–29 yrs (*D25u*)							1.864	19.266	18.052	0.917	7.022	7.474	0.698	5.720	6.328	0.634	3.661	4.887
Unemploy rate 30+ (*D30+U*)							0.578	5.946	9.165	-0.357	-2.724	-2.730	-0.466	-3.761	-3.683	0.369	2.094	4.744

(*): The variable coefficients at time *t* with lagged effects incorporated into the model are interpreted as the long-term effects of the variable on the participation rate. The real minimum wage (*MW*), the generosity of employment insurance (*GEI*) and the unemployment rate (*U*) all have lagged effects on the participation rate. This interpretation of the coefficients derives directly from the algebraic form of the equation. All the continuous dependent and independent variables of the model are expressed in logs.

(**): The variable *z*-asy represents the *t*-statistic constructed from a robust estimator of the covariance matrix of the estimated coefficients. This estimator is convergent in the presence of heteroskedasticity. The variable *z*-asy provides an unbiased asymptotic measure of the test.

Table 9
Estimates of welfare participation rates, Men

Difference-in-difference model, OLS, 77.1 to 98.4, quarterly data
Lag specification: MW= 1, GEI = 4, U =4. (*)

	(A)			(B)			(C)			(D)			(E)			(F)		
Fixed effects by age group	**Yes**			**Yes**			**Yes**			**Yes**			**Yes**			**Yes**		
Seasonal effects by age group	**No**			**Yes**			**Yes**			**Yes**			**Yes**			**Yes**		
Annual effects	**No**			**No**			**No**			**Yes**			**Yes**			**Yes**		
Annual effects by age group	**No**			**No**			**No**			**No**			**No**			**Yes**		
Effect of the 90-93 recession	**No**			**No**			**No**			**No**			**Yes**			**Yes**		
Effect of other factors (unemployment rate, benefits schedule, etc.)	**No**			**No**			**Yes**			**Yes**			**Yes**			**Yes**		
Fisher Test (confidence level = 5%). Each test is sequential: (B) with respect to (A), (C) with respect to (B), etc.				• F = 0,298 ➔ Seasonal effect not significant.			• F =93,428 ➔ Exogenous factors significant.			• F =53,754 ➔ Time effect significant.			• F =5,063➔ Recession effect significant.			F =6,024 ➔ Age*time effect significant.		
	Coeff	***t*-ols**	***z*-asy (**)**	**Coeff**	***t*-ols**	***z*-asy**	Coeff	*t*-ols	*z*-asy	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**
Constant	-4.171	-81.933	-94.367	-4.162	-49.495	-64.342	-8.967	-3.038	-5.727	-4.631	-2.174	-1.965	-1.338	-0.646	-0.793	-1.160	-0.333	-0.643
Common Post Effect (*Dp*)	0.401	5.178	8.352	0.402	5.125	8.412	0.049	0.523	1.006	0.139	1.886	1.349	0.215	2.944	3.867	-0.060	-0.640	-1.505
Post 18 years effect (*D18Dp*)	-0.565	-5.160	-6.531	-0.564	-5.079	-6.565	-1.108	-6.343	-5.924	-0.717	-4.569	-2.792	-0.548	-3.629	-4.676	-0.157	-1.213	-2.424
Post 19 years effect (*D19Dp*)	-0.432	-3.942	-4.919	-0.430	-3.871	-4.931	-0.977	-5.610	-5.272	-0.593	-3.781	-2.317	-0.448	-2.968	-3.965	-0.161	-1.239	-2.513
Post 20 years effect (*D20Dp*)	-0.256	-2.341	-2.916	-0.254	-2.289	-2.914	-0.788	-4.531	-4.269	-0.409	-2.609	-1.601	-0.374	-2.476	-3.360	-0.178	-1.378	-2.797
Post 21 years effect (*D21Dp*)	-0.084	-0.765	-0.956	-0.081	-0.733	-0.936	-0.598	-3.446	-3.231	-0.224	-1.430	-0.878	-0.234	-1.550	-2.092	-0.029	-0.228	-0.447
Post 22 years effect (*D22Dp*)	0.062	0.562	0.698	0.064	0.578	0.735	-0.442	-2.549	-2.354	-0.070	-0.445	-0.272	-0.266	-1.758	-2.430	-0.079	-0.616	-1.179
Post 23 years effect (*D23Dp*)	0.170	1.554	1.910	0.172	1.553	1.953	-0.338	-1.948	-1.770	0.036	0.229	0.139	-0.275	-1.815	-2.539	-0.068	-0.530	-0.963
Post 24 years effect (*D24Dp*)	0.245	2.238	2.754	0.247	2.227	2.802	-0.278	-1.597	-1.452	0.102	0.648	0.394	-0.285	-1.874	-2.694	-0.069	-0.532	-1.001
Post 25–29 years effect (*D2529Dp*)	0.398	3.632	4.448	0.400	3.600	4.503	0.025	0.079	0.047	0.278	1.236	0.665	0.296	1.443	1.502	-0.130	-0.607	-1.177
Real b s 18–24 years (*D1824Bm*)							0.584	3.527	3.060	0.070	0.452	0.283	0.119	0.797	1.017	0.232	2.181	3.327
Real b s 25–29 years (*D2529Bm*)							0.311	0.864	0.511	-0.018	-0.071	-0.039	-0.533	-2.169	-2.095	0.331	1.242	2.217
Real b s 30 years + (*D30+Bm*)							0.709	1.514	2.842	0.496	1.578	1.364	0.308	1.002	1.274	0.092	0.171	0.362
Real minimum wage (*MW*)							-0.670	-11.427	-9.189	-1.236	-3.171	-3.633	-1.968	-5.130	-5.800	-2.048	-8.114	-8.898
EI generosity (*GEI*)							-0.510	-13.376	-13.868	-0.224	-1.816	-2.062	-0.414	-3.372	-3.630	-0.412	-5.125	-6.047
Population rate							0.551	7.673	5.830	0.386	7.564	4.948	0.672	12.793	9.239	0.202	0.787	0.754
Unemploy rate 18 years (*D18u*)							2.276	21.562	19.727	1.145	8.228	7.507	1.184	9.213	8.507	1.103	5.791	3.632
Unemploy rate 19 years (*D19u*)							2.383	22.764	25.244	1.263	9.084	9.715	1.271	9.906	11.057	1.083	5.672	5.541
Unemploy rate 20 years (*D20u*)							2.373	22.821	28.606	1.264	9.099	10.331	1.210	9.436	10.966	1.097	5.761	4.434
Unemploy rate 21 years (*D21u*)							2.323	22.435	29.366	1.225	8.812	10.104	1.133	8.838	10.377	1.100	5.786	6.076
Unemploy rate 22 years (*D22u*)							2.272	21.970	24.733	1.178	8.477	9.597	1.012	7.891	9.038	0.978	5.138	5.697
Unemploy rate 23 years (*D23u*)							2.251	21.781	21.171	1.160	8.348	9.277	0.955	7.441	8.356	0.880	4.621	4.765
Unemploy rate 24 years (*D24u*)							2.218	21.465	19.197	1.129	8.122	8.696	0.902	7.017	7.738	0.837	4.404	5.904
Unemploy rate 25–29 yrs (*D25u*)							2.134	20.450	19.616	1.057	7.575	8.381	0.821	6.376	7.299	0.745	3.882	5.020
Unemployment rate 30+ (*D30+U*)							1.075	10.241	13.653	0.002	0.013	0.013	-0.109	-0.832	-0.821	0.433	2.221	5.236

(*): The variable coefficients at time t with lagged effects incorporated into the model are interpreted as the long-term effects of the variable on the participation rate. The real minimum wage (MW), the generosity of employment insurance (GEI) and the unemployment rate (U) all have lagged effects on the participation rate. This interpretation of the coefficients derives directly from the algebraic form of the equation. All the continuous dependent and independent variables of the model are expressed in logs.

(**): The variable z-asy represents the t-statistic constructed from a robust estimator of the covariance matrix of the estimated coefficients. This estimator is convergent in the presence of heteroskedasticity. The variable z-asy provides an unbiased asymptotic measure of the test.

Table 10
Estimates of welfare participation rates, Women

Difference-in-difference model, OLS, 77.1 to 98.4, quarterly data
Lag specification: MW= 1, GEI = 4, U =4. (*)

	(A)			(B)			(C)			(D)			(E)			(F)		
Fixed effects by age group	**Yes**			**Yes**			**Yes**			**Yes**			**Yes**			**Yes**		
Seasonal effects by age group	**No**			**Yes**			**Yes**			**Yes**			**Yes**			**Yes**		
Annual effects	**No**			**No**			**No**			**Yes**			**Yes**			**Yes**		
Annual effects by age group	**No**			**No**			**No**			**No**			**No**			**Yes**		
Effect of the 90-93 recession	**No**			**No**			**No**			**No**			**Yes**			**Yes**		
Effect of other factors (unemployment rate, benefits schedule, etc.)	**No**			**No**			**Yes**			**Yes**			**Yes**			**Yes**		
Fisher Test (confidence level = 5%). Each test is sequential: (B) with respect to (A), (C) with respect to (B), etc.				• F = 0,060 → Seasonal effect not significant,			• F =69,383 → Exogenous factors significant,			• F =44,796 → Time effect significant,			• F =2,368 → 16. Recession effect significant,			F =9,063 → Age*time effect significant,		
	Coeff	***t*-ols**	***z*-asy (**)**	**Coeff**	***t*-ols**	***z*-asy**	Coeff	*t*-ols	*z*-asy	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**	**Coeff**	***t*-ols**	***z*-asy**
Constant	-4,225	-107.267	-273.668	-4.224	-64.660	-187.616	-16.457	-4.592	-9.981	-17.187	-6.725	-6.388	-16.123	-5.944	-5.870	-2.342	-0.723	-1.119
Common Post Effect (Dp)	0.024	0.396	1.327	0.024	0.389	1.330	0.213	1.907	4.239	0.550	6.552	6.033	0.598	6.870	7.010	0.101	1.116	1.804
Post 18 years effect ($D18Dp$)	-0.389	-4.584	-6.914	-0.389	-4.509	-6.944	-0.738	-4.115	-4.671	-0.889	-5.785	-4.609	-0.898	-5.580	-7.075	-0.208	-1.748	-3.002
Post 19 years effect ($D19Dp$)	-0.346	-4.077	-5.863	-0.345	-3.999	-5.874	-0.705	-3.943	-4.473	-0.863	-5.618	-4.492	-0.824	-5.121	-6.579	-0.208	-1.745	-3.023
Post 20 years effect ($D20Dp$)	-0.224	-2.645	-3.612	-0.224	-2.593	-3.617	-0.583	-3.267	-3.702	-0.746	-4.863	-3.890	-0.706	-4.390	-5.678	-0.191	-1.606	-2.806
Post 21 years effect ($D21Dp$)	-0.079	-0.938	-1.259	-0.079	-0.913	-1.254	-0.425	-2.385	-2.701	-0.593	-3.866	-3.091	-0.692	-4.303	-5.576	-0.211	-1.771	-2.999
Post 22 years effect ($D22Dp$)	0.070	0.823	1.093	0.071	0.821	1.114	-0.269	-1.512	-1.706	-0.439	-2.864	-2.286	-0.579	-3.593	-4.615	-0.118	-0.992	-1.687
Post 23 years effect ($D23Dp$)	0.188	2.220	3.055	0.189	2.191	3.079	-0.142	-0.796	-0.885	-0.310	-2.023	-1.603	-0.560	-3.472	-4.501	-0.118	-0.997	-1.625
Post 24 years effect ($D24Dp$)	0.291	3.427	4.876	0.291	3.379	4.912	-0.043	-0.243	-0.267	-0.206	-1.345	-1.063	-0.555	-3.429	-4.538	-0.128	-1.075	-1.767
Post 25–29 years effect ($D2529Dp$)	0.450	5.309	8.375	0.451	5.226	8.424	-0.507	-1.604	-1.549	-0.790	-3.405	-3.630	-0.754	-3.258	-4.265	-0.165	-0.860	-1.850
Real b s 18–24 years ($D1824Bm$)							-0.084	-0.526	-0.508	-0.047	-0.320	-0.252	0.060	0.388	0.558	0.078	0.828	1.287
Real b s 25–29 years ($D2529Bm$)							0.654	1.793	1.651	0.870	3.302	3.675	0.516	1.813	2.506	0.195	0.817	1.860
Real b s 30 years + ($D30{+}Bm$)							2.221	3.902	8.316	2.523	6.525	6.278	2.522	6.128	6.230	0.021	0.041	0.066
Real minimum wage (MW)							-0.686	-13.207	-11.968	-0.661	-1.813	-2.003	-0.989	-2.551	-2.848	-1.058	-4.816	-5.447
EI generosity (GEI)							-0.295	-9.312	-10.729	-0.094	-0.817	-0.908	-0.197	-1.584	-1.711	-0.196	-2.809	-3.153
Population rate							0.600	9.368	7.769	0.424	9.028	7.305	0.642	12.245	10.082	0.355	1.594	1.422
Unemploy rate 18 years ($D18u$)							1.621	17.408	17.099	0.794	6.120	5.700	0.811	6.264	5.904	0.953	5.764	4.310
Unemploy rate 19 years ($D19u$)							1.772	19.186	22.402	0.957	7.382	7.604	0.972	7.520	8.039	0.874	5.275	4.858
Unemploy rate 20 years ($D20u$)							1.837	20.013	22.761	1.034	7.978	8.709	1.028	7.967	8.864	0.906	5.477	4.287
Unemploy rate 21 years ($D21u$)							1.807	19.768	22.855	1.015	7.835	8.836	0.948	7.348	8.247	0.718	4.345	3.984
Unemploy rate 22 years ($D22u$)							1.784	19.539	21.883	0.997	7.697	8.718	0.907	7.024	7.881	0.512	3.098	3.545
Unemploy rate 23 years ($D23u$)							1.680	18.411	18.931	0.897	6.920	7.600	0.766	5.928	6.521	0.655	3.962	4.292
Unemploy rate 24 years ($D24u$)							1.565	17.152	16.262	0.784	6.048	6.314	0.620	4.798	5.145	0.545	3.304	5.374
Unemploy rate 25–29 yrs ($D25u$)							1.236	13.419	14.111	0.469	3.608	3.727	0.320	2.466	2.621	0.373	2.234	3.765
Unemploy rate 30+ ($D30{+}U$)							-0.010	-0.106	-0.164	-0.765	-5.849	-5.752	-0.849	-6.449	-6.201	0.248	1.456	3.076

(*): The variable coefficients at time t with lagged effects incorporated into the model are interpreted as the long-term effects of the variable on the participation rate. The real minimum wage (MW), the generosity of employment insurance (GEI) and the unemployment rate (U) all have lagged effects on the participation rate. This interpretation of the coefficients derives directly from the algebraic form of the equation. All the continuous dependent and independent variables of the model are expressed in logs.

(**): The variable z-asy represents the t-statistic constructed from a robust estimator of the covariance matrix of the estimated coefficients. This estimator is convergent in the presence of heteroskedasticity. The variable z-asy provides an unbiased asymptotic measure of the test.

Table 11
Test of stationarity

Variables with time effects,
Individual Dickey-Fuller test (adjusted)

	Case 1			Case 2		
Variables	True Process (H_o accepted): $y_t = \sum_{i=1}^{l} \varsigma_i \Delta y_{t-i} + \alpha + y_{t-1} + \varepsilon_t$ Estimated Process: $\Delta y_t = \sum_{i=1}^{l} \varsigma_i \Delta y_{t-i} + \alpha + (\rho - 1) y_{t-1} + \delta t + \varepsilon_t$			True Process (H_o accepted) : $y_t = \sum_{i=1}^{l} \varsigma_i \Delta y_{t-i} + \alpha + y_{t-1} + \varepsilon_t$ Estimated Process: $\Delta y_t = \sum_{i=1}^{l} \varsigma_i \Delta y_{t-i} + \alpha + (\rho - 1) y_{t-1} + \varepsilon_t$		
	# obs.	l	$(\rho - 1)$	*# obs.*	l	$(\rho - 1)$
Unemployment rate (*U*)	86	4	-0.152	87	4	-0.109
Real minimum wage (*MW*)	87	4	-0.017	87	4	-0.034
Nonpecuniary gen. of E.I. (*GEI*)	87	4	-0.130	87	4	-0.032
Real benefits schedule <30 years	87	4	-0.074	87	4	-0.018
Real benefits schedule $\geq$ 30 years	87	4	-0.132	87	4	-0.020

Note : All variables are expressed in logs. The data is quarterly and the observations extend from the first quarter of 1976 until the last quarter of 1998, the exact period varying with the lag *l* of the variable. *: The hypothesis of nonstationarity (H_o) rejected at 5%. The critical values are from Hamilton (1994). Notice that the critical values for case two derive from the standard normal distribution [cf. Hamilton (1994)].

Pooled variables
Levin et Lin test (1993)

	Case 1						Case 2					
Variables	True Process (H_o accepted): $y_{it} = \sum_{j=1}^{l} \varsigma_{ij} \Delta y_{i,t-j} + \alpha_i + y_{i,t-1} + \varepsilon_{it}$ Estimated Process: $\Delta y_{it} = \sum_{j=1}^{l} \varsigma_{ij} \Delta y_{i,t-j} + \alpha_i + (\rho_i - 1) y_{i,t-1} + \delta_i t + \varepsilon_{it}$						True Process (H_o accepted) : $y_{it} = \sum_{j=1}^{l} \varsigma_j \Delta y_{i,t-j} + y_{i,t-1} + \varepsilon_{it}$ Estimated Process: $\Delta y_{it} = \sum_{j=1}^{l} \varsigma_{ij} \Delta y_{i,t-j} + \alpha_i + (\rho_i - 1) y_{i,t-1} + \varepsilon_{it}$					
	# obs.	l	$t_{\rho-1}$	Adjustment factors		Adjusted $t_{\rho-1}$	*# obs.*	l	$t_{\rho-1}$	Adjustment factors		Adjust. $t_{\rho-1}$
				Mean	Standard error					Mean	Standard error	
Participation rate	9*85	5	-0,462	-0,578	0,728	0,396	9*85	5	-1,728	-0,521	0,789	-1,217

Notes: The statistic $t_{\rho-1}$ follows a normal standard distribution. *: The hypothesis of nonstationarity (H_o) rejected at 5%.

Table 12
Pedroni Test of cointegration

	Estimated equation: $y_{it} = \alpha_0 + \alpha_1^{'} D_T + \alpha_2 d_P + \alpha_3^{'} D_T d_P + \alpha_4^{'} x + D_T^{'} \text{A} x + \mu_{it}$ (See equation 4, p.11, for a description of the parameters and variables)						
Tests (group type)	# *obs.*	*l*	m	Statistics	Adjustment Factors		Adjusted Statistics
					Mean	Variance	
ρ -test	9*85	5	7	-174,487	-36,494	140,756	-5,479 *
t -test (non- parametric)	9*85	5	7	-18,794	-4,217	0,518	-8,535 *
t -test (parametric)	9*85	5	7	-22,596	-4,217	-4,217	-13,818 *

Notes: * The hypothesis of non-cointegration (H_o) is rejected at 5%. The adjusted statistics all follow the standard normal distribution. The adjustment factors are drawn from Pedroni (1999). *l* designates the degree of the lag in the calculation of the of the statistics, and *m* the number of regressors in the estimated equation; the lagged explanatory variables are not incorporated in the equation, nor are the interactive effects between the cohort-specific effects and the time effect.